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An Empirical Investigation of the Nexus among Money Balances, Commodity Prices and Consumer Goods' Prices

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Abstract

This paper aims to identify the nexus between the excess of liquidity in the United States and commodity prices over the 1983–2006 period. In particular, it assesses whether commodity prices react more powerfully than consumer goods' prices to changes in real money balances. Within a cointegrated vector autoregressive framework, the author investigates whether consumer prices and commodity prices react to excess liquidity, and if the different price elasticities of supply for goods and commodities allow for differences in the dynamic paths of price adjustment to a liquidity shock. The results show a positive relationship between real money and real commodity prices and provide empirical evidence for a stronger response of commodity prices with respect to consumer goods' prices. This could imply that, if the magnitude of the reaction is due the fact that consumer goods' prices are slower to react, then their long-run value can be predicted with the help of commodity prices. The findings support the view that the latter should be considered as a valid monetary indicator.

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An Empirical Investigation of the Nexus among Money Balances, Commodity Prices and Consumer Goods' Prices

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1. Introduction

In the last four decades, the volatility of commodity prices generated turbulence in the global economy, affecting importing and exporting countries in opposite and vigorous ways. Nonetheless, the attention of the literature to the topic seemed to be proportional to price growth, declining in relatively tranquil periods and rising when the commodities prices were back at high levels. The huge variations of the last decade generated a renewed interest in the topic.

In the 1970s, the popular view was that commodity prices were defined as a result of evolution in the relevant commodity market, playing an important role in the stagflation of that decade. However, this idea has been strongly challenged. An increase in the expected inflation rate due, for example, to an increase in money supply, causes agents to shift from money to commodities, provoking a rise in prices. Therefore, increases in the price of oil and other commodities could be the result of an exceedingly expansionary policy, rather than an exogenous inflationary supply shock.

Falling commodity prices in the 1980s and 1990s were not considered as interesting as rising prices, even though oil producers such as Mexico and Russia were experiencing important revenue losses and countries such as Argentina, Brazil and Australia were suffering from low agricultural prices.

After collapsing in the second half of 2008, commodity prices stabilized in early 2009 and subsequently staged a comeback. Such behavior is in contrast with what happened during past recessions. In previous global downturns, prices typically continued to fall into the early phases of recovery or rose at rates far below the increases recorded in recent months. An exception is the price of oil, which recorded meaningful increases early in previous recoveries.

Thus, the recent happenings of quick commodity price increases and higher volatility following the easy monetary stance in the US, matched with similar accommodative policies in the euro area and Japan, led some to infer some causal role for monetary changes in driving commodity prices and ultimately inflation.

Drawing on Dornbusch (1976), the idea of overshooting has been adapted to analyze theoretically the relationship between money, consumer prices and commodity prices by Frankel (1986). The latter argued that tightening monetary policy has relevant effects on commodity prices because they are flexible, whereas other goods' prices are sticky. Thus, commodity prices overshoot their new equilibrium in the short-run in order to generate an expectation of future appreciation sufficient to offset the higher interest rate.

This paper aims at identifying the nexus between the excess of liquidity in the United States and commodity prices over the 1983-2006 period. In particular, it tests whether the latter react more powerfully than consumer goods' prices to changes in real money balances.

Within a cointegrated vector autoregressive (VAR) framework, it is investigated whether consumer prices and commodity prices react to excess liquidity in the US, and if the different price elasticities of supply for goods and commodities allow for differences in the dynamic paths of price adjustment to a liquidity shock.

The results show a positive relationship between real money and real commodity prices and provide empirical evidence for a stronger response of commodity prices with respect to consumer goods' prices. This could imply that, if the magnitude of the reaction is due to the fact that consumer goods' prices are slower to react, then their long-run value can be predicted with the help of the commodity prices.

The structure of the paper is as follows. In section 2, some contributions on the issue are presented. Section 3 describes the hypothesis to be tested. In section 4 the empirical strategy, tests results and interpretation of the findings are presented. Section 5 ends the paper reporting the conclusions.

2. A review of the literature

Before discussing the empirical analysis employed, some of the main contributions on the relationship between monetary policy, consumer prices and commodity prices will be reviewed.

Drawing on Dornbusch's exchange rate overshooting model, Frankel (1986) provides a theoretical framework to analyze the impact of money supply shocks on commodity prices. He substitutes the price of agricultural goods for the price of foreign exchange and argues that the reason for the overshooting phenomenon is that prices for agricultural and mineral products adjust rapidly, while the price of goods adjusts more slowly. In fact, the hypothesis is that commodities are exchanged in a reactive auction market in which the supply cannot be easily expanded, whereas the consumer goods market enjoys a copious supply.

Frankel illustrates the dynamic starting from a monetary contraction that is expected to be permanent and that would eventually lead to an equal fall of consumer and commodity goods' prices in the absence of other disturbances. However, given the manufactured price stickiness in the short-run the nominal money supply contraction is a reduction in the real money supply. Such a reduction should be offset by an interest rate rise. Nevertheless, the arbitrage condition implies that, since commodities are storable, the interest rate cannot grow more than the expected rate of increase in the commodity prices plus the storage cost. Therefore, the spot price of commodities must fall today and must do so until the moment in which there is an expectation of future appreciation that is sufficient to offset the higher interest rate.

Over the last three decades the role of commodity prices in setting monetary policy has been faced by many researchers. On the one hand, it has been argued that commodity prices may be an earlier indicator of the current state of the economy because, as assumed by Frenkel (1986) these prices are usually set in continuous auction markets with efficient information (see Olivera, 1970; Garner, 1989; Marquis and Cunningham, 1990; Cody and Mills, 1991). Thus, some policymakers became advocates of using commodity prices as a leading indicator of inflation and endorsed policy proposals using commodity prices as a guide to adjust short run money growth target ranges (see Garner, 1989). A rise in commodity prices may indicate to policymakers that the economy is growing too rapidly and hence inflation is inclined to rise. In such a case, the monetary authority may observe the rising commodity prices and respond by raising interest rates to tighten money supply.

On the other hand, the counterargument is that commodity prices cannot be used effectively in formulating monetary policy because they are subject to large, market-specific shocks, which may not have macroeconomic implications (see Marquis and Cunningham, 1990; Cody and Mills, 1991). However, many others (see Bessler, 1984; Pindyck and Rotemberg, 1990; Hua, 1998) argue that commodity price movements are the result of monetary or macroeconomic changes and that the causality should run from macroeconomic/monetary variables to commodity prices. Barsky and Kilian (2002) offer another important contribution. They argue that monetary expansions and contractions could generate stagflation of important magnitudes, by providing evidence about the role of monetary fluctuations in determining the prices of oil and, in particular, the prices of industrial commodities that preceded the 1973 oil price increase. Bernanke et al. (1997) investigate the relationship between the oil price shocks, US monetary policy and the business cycle. They assert that a relevant part of the effect of oil price shocks on the economy comes from tighter monetary policy resulting from the change in oil prices and not from the change in oil prices *per se*.

Another piece of literature analyzes the impact of the commodity price evolutions on the behavior of monetary policy and its informational role for formulating it. Awokuse and Yang (2003) argue that commodity price indicators contain important information about the future movements of macroeconomic variables. Bhar and Hamori (2008) assess the information content of commodity prices for monetary policy. Using a cross correlation approach between economic activity and commodity futures prices, they affirm that commodity prices can serve as suitable information for monetary policy.

Fuhrer and Moore (1992) investigate the relationships between asset prices and inflation in a Keynesian model in which monetary policy controls inflation by manipulating the federal funds rate. They find that the indicator properties of asset prices are quite sensitive to the monetary policy rule. Hamori (2007) empirically analyzes the relationship between the commodity prices index and macroeconomic variables in Japan, arguing that the former and the general price level are closely related, with movements in commodity prices leading movements in the general price level. However, he specifies that the commodity price index was found to be valid as a leading indicator of the consumer price index before the zero interest policy was introduced, as afterwards the relationship ceased to exist.

Other studies, such as Surrey (1989), Boughton and Branson (1990, 1991), and Browne and Cronin (2007), investigate empirically the potential importance of monetary conditions on the relationship between commodity prices and consumer goods' prices. However, they all use different empirical techniques or different specifications from the one employed in this paper.

3. A Model of Price Dynamics

This section presents the theoretical framework through which the investigation aims to answer the following research question: do commodity prices react more strongly than consumer goods' prices?

As mentioned, the commodity price overshooting theory was advanced by Frankel (1986). The essence of this theoretical framework is that the short-term reaction to an expansionary monetary policy produces an overshoot of the commodity prices and a more delayed reaction in the consumer goods market. In the long-run, consumer prices adjust to the new equilibrium. However, Frenkel's theory is fundamentally based on the assumption that commodity prices react more strongly in the short-run than the consumer goods' prices.

In order to test the hypothesis it is allowed for a two-good economy, and commodities and consumer goods with prices p^{COM} and p^{CPI}, respectively. The substantial difference between these two goods is that consumer prices are decided in a market with a supply that adjusts to the changes in demand, while the commodity prices are decided in a market restricted in supply and with high transaction costs, due to transportation expenses. Therefore, consumer goods' prices are sticky, whereas commodity prices are not.

The rationale for this assumption can be found in the current scenario where many low-cost producers (especially in developing countries) are generating additional supply of consumer goods, while commodity supply is constrained by natural factors. Furthermore, the speed of the adjustment depends on the fact that participants in the commodity markets are usually more equally informed than their consumer goods' counterparts.

Graphically, the two markets can be represented as in Figure 1.

[Figure 1 about here]

Figure 1 shows the price-quantity changes as a result of a monetary expansion in markets with high (left graph) and low (right graph) price elasticity of supply. The aggregated supply of price elastic goods in the short-run is characterized by infinite price elasticity so that additional demand brought about by a liquidity shock (from D to D') can be satisfied without any price increase. Consequently, the liquidity shock translates into an increase in output achieving a new short-run equilibrium at p'^{CPI}. In contrast, goods characterized by restrictions in supply, cannot be expanded easily and are thus quantity-insensitive to a monetary expansion. Additional demand is then fully reflected in a rise in commodity prices.

In the long-run, prices will also react on the price elastic goods market if the well-documented neutrality of money holds; any change in money supply is met with a proportional change in the price level that keeps real money and real output in both markets unchanged.

More formally, the general price level is:

$$p = \lambda p^{COM} + (1 - \lambda)p^{CPI}, 0 < \lambda < 1 \tag{1}$$

A once-off increase of μ percent in the money supply in period t produces an increase in the general price level by $(1 + \mu_t)p_{t-1}$. However, given the initial assumption, such increase fully translates into the commodity price p^{COM}, because p^{CPI} is sticky. Thus, the price relationship at time t will be:

$$p_t = \lambda p_t^{COM} + (1 - \lambda)p_{t-1}^{CPI} \tag{2}$$

Assuming no further changes in the money supply in period $t + 1$, the general level of prices at time $t + 1$ will be the same as in t.

$$p_{t+1} = \lambda p_{t+1}^{COM} + (1 - \lambda)p_{t+1}^{CPI} \tag{3}$$

This allows setting the right hand side of Equation 2 to be equal to the right end side of Equation 3.

$$\lambda p_t^{COM} + (1 - \lambda)p_{t-1}^{CPI} = \lambda p_{t+1}^{COM} + (1 - \lambda)p_{t+1}^{CPI} \tag{4}$$

As $p_{t-1}^{CPI} = p_t^{CPI}$ and $p_{t+1}^{COM} = (1 + \mu_t)p_{t-1}^{COM}$, after some algebra the following equation is obtained:

$$p_{t+1}^{CPI} - p_t^{CPI} = [\lambda(1 - \lambda)][p_t^{COM} - (1 + \mu_t)p_{t-1}^{COM}] \tag{5}$$

The different dynamics of the adjustment processes implies that the size of the change of the consumer goods' prices in period t can be predicted by observing the spread between the current period price of the commodities (p_t^{COM}) and the equilibrium value to which it must adjust in period $t + 1$, ($(1 + \mu_t)p_{t-1}^{COM}$), which is dependent on the monetary shock in the current period (μ_t).

4. Empirical Analysis

First, this section documents the strategy chosen for the analysis and the data used and describes the empirical analysis. Secondly, the results from the estimation of the model are presented and discussed.

4.1. Empirical Strategy

The analysis is carried out using a cointegrated VAR model. Formally the model is six dimensional VAR with independent and identically-distributed Gaussian errors:

$$X_t = \mathrm{A}_1 X_{t-1} + \cdots + \mathrm{A}_k X_{t-k} + \Phi D_t + \varepsilon_t,\ t = 1, \ldots T \tag{6}$$

where X_t is the following vector of variables:

$$X_t = \begin{bmatrix} y^r \\ m^r \\ i_m \\ i_b \\ \Delta P \\ CRB - CPI \end{bmatrix} \tag{7}$$

Namely real output, real money, short term and long term interest rates, inflation and real commodity prices, and D_t is a vector of deterministic components.

The Vector Error Correction Model (VECM) representation of the VAR model combines levels and differences as follows:

$$X_t = \Pi X_{t-1} + \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1} \Delta X_{t-k-1} + \Phi D_t + \varepsilon_t,\ t = 1, \ldots T \tag{8}$$

where Π are the effects in the long-run and Γ_1 contains the short-run information. The inconsistency with the X_t integrated of order one is solved by transforming the multivariate model and reducing the rank of Π to $r < p$ with p being the number of variables. The reduced rank matrix can be factorized into two rxp matrices α and β ($\Pi = \alpha\beta'$). The factorization provides r stationary linear combinations of the variables called cointegrating vectors, and $p - r$ common stochastic trends of the system.

Within a cointegrated VAR framework, the common shocks, or common stochastic trends, are hitting all the variables simultaneously. Since an impulse response analysis implies assuming a certain structure of shocks, namely that one variable is exclusively hitting the variable we are interested in, this has been neglected as such assumption cannot be tested (see Juselius, 2006).

4.2. Data and Unit Root Tests

The choice of the country has to do with the fact that, even if the US accounts for less than one-third of world GDP, its importance in the monetary and financial system is evidently higher than that. The period under analysis goes from the first quarter of 1983, to the first quarter of 2008. Such span corresponds to a fairly stable period in terms of inflation growth.

As an indicator of the nominal money supply, data on the M2 aggregate have been downloaded from the International Financial Statistics (IFS) dataset, as well as data on nominal gross domestic product (GDP), consumer price index (CPI), 3-month Treasury bill rate and 10-years government bond yield. As a proxy for the commodity prices the Commodity Research Bureau (CRB) index has been adopted. This index measures the combined movements in the prices of 22 basic commodities whose markets are assumed to be among the first to be influenced by changes in economic conditions and therefore by the monetary policy.

In order to perform an I(1) analysis, a nominal to real transformation has been performed. Since nominal GDP and nominal money supply show clear features of I(2) processes, they have been considered in real terms. Likewise, some transformations have been done on price indexes. On the one hand, inflation has been considered instead of the CPI index (typically I(2)); on the other hand, the difference between the CRB index and the CPI index has been taken (real commodity prices hereafter).

Real money, real GDP, inflation and real commodity prices have been taken in logs. The interest rates have been transformed into quarterly rates. Table 1 presents some descriptive statistics of the variables considered[2].

[Table1 about here]

The unit root properties of the series are tentatively investigated using two unit root tests. The first unit root test performed is the Augmented Dickey-Fuller (ADF) test, for which stationarity serves as the null hypothesis. However, it should be noted that the ADF test could fail to distinguish between a unit root and a near unit root process and it can happen that indicates that a series contains a unit root when it does not (Perron 1989). Thus, a second unit root test is adopted, namely the Kwiatkowsk, Phillips, Schmidt and Shin (KPSS) test, for which the null hypothesis is non-stationarity.

[Table 2 about here]

As shown in Table 2, the results from both the tests employed confirm that all the variables in the system are integrated of order one at 5% significance level. Nonetheless, it should be remembered that such tests are not reliable in the presence of breaks or shifts in the series.

Figure 2 and 3 show the graphs of the series in levels and first differences.

[Figure 2 about here]

[Figure 3 about here]

A graphical inspection of the data reveals that the assumption of constant mean does not seem appropriate for the level of the variables, whereas it does much more so for the differenced series. The assumption of constant variance seems to be approximately satisfied for the differences, even though some variability can be observed in most of the series. In order to account for the biggest departures from assumptions, three permanent blip dummies have been added to the model in correspondence of the biggest residuals. The first one takes the value one during the fourth quarter of 1984, when Ronald Reagan was re-elected as President of the US. The second one controls for the dramatic slump in inflation during the

[2] Note that for both the CPI and CRB index the value 100 is given to the first quarter of 1983.

second quarter of 2006, probably due to an unexpected plunge in energy prices. The last dummy controls for the economic impacts of the terrorists attack of September 11, 2001.

Moreover, it is worth noting that the graph of the real money shows two breaks in trend. A constant increase in the variable is observed from the beginning of the sample until 1987. A period of approximately zero growth follows and lasts until 1995, when it starts increasing as fast as in the first period. Thus, the slope of the two growing trends seems to be roughly the same. Even though the introduction of two breaks has been attempted, it turned out more satisfactory to allow for trends in the levels. This should have the effect of averaging out the aforementioned breaks. Moreover, since it is not possible to know *a priori* whether these trends cancel out in the cointegrating relations, the chosen specification allows these to be trend-stationary and have non-zero intercepts.

4.3. Lag Length Selection and Residual Analysis

Table 3 shows that the Schwartz criterion (SC) suggests $k = 1$ and the Hannan-Quinn (HQ) criterion suggests $k = 2$. However, when imposing $k = 1$ the other misspecifications tests become much worse, implying that the SC might have penalized too much. The LM tests in the last two columns show the left-over residual autocorrelation in each VAR(k) model and seem to accept the absence of autocorrelation for the VAR with 2 lags.

[Table 3 about here]

Table 4 presents the results of the multivariate residual analysis. In particular the null hypothesis of no autocorrelation is rejected at both the first and the second lag. Given the relatively small sample, it is not advisable to rely on the asymptotic properties of the estimator, thus the normality assumption turns out to be relevant and it is safely accepted. Lastly the null hypothesis of no ARCH effects is accepted at both lags.

[Table 4 about here]

Table 5 shows the results of the univariate residual analysis. The output of the ARCH and Normality tests reflect the good results of the multivariate analysis. The skewness and kurtosis statistics are close to the normal distribution values, suggesting that by inserting the three blip dummies the biggest outliers should have been controlled for.

[Table 5 about here]

Overall, both the multivariate and univariate tests suggest that the residuals are well behaved and therefore that the model is well specified.

4.4. Rank Determination

The cointegration rank is determined according to Johansen (1996) LR Trace test. When the sample is small, the asymptotic distributions are generally poor approximations to the true distributions. To secure a correct test size one can apply the small sample Bartlett corrections developed in Johansen (2002). The asymptotic distribution of the rank test statistic differs depending on the deterministic components in the model and on almost any type of dummy variable[3]. Therefore, the safest procedure is to simulate the new critical values. Table 6 presents the Trace test results with the Bartlett corrections and the simulated critical values.

[Table 6 about here]

The choice of the cointegration rank is not clearly defined, in fact the Trace test Bartlett corrected suggests that $r = 3$ is accepted only at 10% significance level. Looking at the significance of the α coefficients of the third cointegration vector in Table 7 it seems that information regarding the real GDP and the real commodity prices would be neglected by choosing $r = 2$.

[Table 7 about here]

A graphical inspection of the cointegrating relations in Figure 4 reveals some symptom of I(2)ness. The first two cointegration relationships look fairly stationary, but the third one presents some indication of cyclical swings. However, it should be observed that the lower panel ($\beta_r' R_{1t}$) corrected for short-run effects of each graph is similar to the upper panel ($\beta_r' x_t$), confirming that the I(2) problem could have been limited (see appendix for a formal I(2) rank test).

[Figure 4 about here]

Moreover, an examination of the characteristic roots shows that the largest unrestricted root for $r = 2$ is 0.90 and for $r = 3$ is 0.78 (pretty far from the unit circle). This seems to confirm the presence of three common stochastic trends. Figure 5 shows the roots of the companion matrix for $r = 3$.

[Figure 5 about here]

4.5. Recursive Tests

The graphs of the recursively calculated fluctuation tests in Figure 6 show that the X-form of $\hat{\tau}_1$ and $\hat{\tau}_3$ are in the rejection region at the beginning of 1995, when the recursion starts. The test statistics remain at a fairly high level until approximately 1998. The recursive graphs of the $\hat{\tau}_2$ suggest that the parameter of the second cointegration relation are considerably constant over the sample period. The overall test in the lower right-hand side

[3] An exception to this are the centered seasonal dummies, which, by construction, sum to zero over time, and hence do not change the asymptotic distribution of the rank tests.

panel picks up the non-constancy at the beginning of the recursions. However, it should be noted that the R-form looks stable in all $\hat{\tau}$, meaning that the instability is only in the short-run coefficients. In general, the eigenvalue fluctuation tests provide a fair picture.

[Figure 6 about here]

The Max test of β Constancy is always lower than one and shows a slightly higher volatility after 2003. Figure 7 confirms what is suggested by the eigenvalue fluctuation tests, namely that the changes in the eigenvalues are due to changes in α.

[Table 7 about here]

4.6. Long-Run Exclusion

The Π matrix gives tentative evidence of long-run exclusion for the variables in the system. If a variable is excludable, the coefficients in the columns must be insignificant. From the PI matrix there are no clear signs that any of the variables can be excluded from the cointegration relations.

A formal LR-test for variable exclusion has been performed. Based on the results in Table 8, it is not possible to exclude any variable at 10% significance level for $r = 3$. However, the real commodity prices seem to be a border-line case.

[Table 8 about here]

4.7. Weak Exogeneity and Pure Adjustment

The Π matrix gives preliminary evidence of weak exogeneity. If a variable is weakly exogenous, the coefficients in the rows must be insignificant; in other words, it represents a pushing force. The only variable that seems not to react to any other variables is the real money.

The formal LR-test for weak exogeneity in Table 9 shows that both real GDP and real money are weakly exogenous at 10% significance level, even though the joint exogeneity is rejected. It has been decided to carry out the analysis including them in the system in order to analyze their impact in the Γ matrix.

[Table 9 about here]

Another LR-test is carried out to test for unit vector in the α matrix. In other words, it tests whether the cumulated disturbances from the i^{th} variable do not enter the common trends (pure adjustment hypothesis). At 10% significance level Table 10 shows that the null hypothesis cannot be rejected for the long term interest rate and the real commodity prices. Thus, these can be considered pulling forces.

[Table 10 about here]

4.8. Identification and Interpretation of the Results

The identifications process starts from the long-run relationships. It is carried out by imposing $r-1$ non-testable[4] just-identifying restrictions on each beta vector of the unrestricted reduced VAR model with rank $r=3$, with no changes to the likelihood function with respect to the under identified model. Table 11 presents the estimation results that seem to suggest that the first beta vector is describing a money demand relationship with an unconventional sign on the inflation variable. The second cointegrating relationship illustrates the expected correlation between real output, real money, real commodity prices and a trend. Finally, the third stationary vector is expressing a relationship between inflation and the interest rate spread.

[Table 11 about here]

The over-identified structure is modeled imposing restrictions accordingly to the t-value of the β vectors' coefficients of the just-identified structure. This is accepted with a fairly large p-value of 0.642 (meaning that the stationarity of the long-run relations cannot be jointly rejected).

Table 12 permits the detection of the pulling forces for each cointegration relation. The money demand relationship is corrected by changes in inflation, whereas the real commodity prices are the only equilibrium correcting force for the second cointegrating relation. Lastly, deviations from the inflation expectations relationship are corrected by inflation itself.

[Table 12 about here]

The short-run identification has been carried out by removing all the non-significant variables from each equation of the cointegrated VAR[5]. The parsimonious structure cannot be rejected with a p-value of 0.257 and is broadly consistent with the classification into pushing and pulling forces[6].

The equation for the interesting variable is reported in Table 13.

[Table 13 about here]

From the inspection of the Ω matrix in Table 14, it is evident that some residuals are highly correlated (positively between real GDP and real money, negatively between inflation

[4] Since we have imposed $r-1$ restrictions, there are zero degrees of freedom, thus restrictions cannot be tested.
[5] Centered seasonal dummies have been left in the equations for each variable, even if insignificant.
[6] The equilibrium correcting role of the real commodity prices for the second cointegrating relationship confirmed.

and real GDP, negatively between inflation and real money and positively between the two interest rates) and this would suggest a simultaneous specification of the model. Nonetheless, such analysis has not been pursued because of the problems in determining the direction of the causation between the mentioned variables.

[Table 14 about here]

Overall, the results confirm the existence of the hypothesized long-run equilibrium relationship between real money and real commodity prices, with an effect from real output. Moreover, there is evidence that the real commodity prices are the only equilibrium correcting variable.

It can be concluded that in order to restore the long-run equilibrium when there is a real excess (lack) of liquidity, the real commodity prices need to increase (decrease). Such increase (decrease) can be achieved through an increase (decrease) in the commodity prices stronger (smaller) than the increase (decrease) in the consumer goods' prices, generating a larger spread between the two as in Figure 8.

[Figure 8 about here]

The Moving Average (MA) representation of the data for $r = 3$ corresponds to $p - r = 3$ common trends. Since $p - r - 1 = 2$ just-identifying restrictions are imposed on each vector, estimates are not unique and the likelihood function is unchanged. However, the space spanned by $\alpha_{\perp}$ and $\beta_{\perp}$ is uniquely determined, so that the estimated long-run impact matrix C is unique[7]. The normalization has been placed on the variables with the highest residual standard errors, namely real GDP, real money and real commodity prices.

The cumulated empirical shocks to the real GDP have had significant and high negative effects on the real commodity prices. The inverse is true, but the impact is much more moderate. The cumulated shocks to real money have had positive effects on inflation, as well as the cumulated shocks to short-term interest rate on long-term term one and vice versa. Moreover, the cumulated shocks to long term interest rate have had positive effects on inflation. Likewise, the cumulated shocks to inflation have had similar positive effects on both the interest rates. Finally, the cumulated shocks to all the variables have had a positive effect on themselves. It seems that the long-term interest rate is purely adjusting (consistently to the previous findings).

In order to impose the over-identifying restrictions on α, joint weak exogeneity for real money and real GDP is tested keeping fixed the restrictions on the β vectors. Since the hypothesis turns out to be rejected at 10% significance level, a zero row in α is imposed only for the real money, for which the null hypothesis cannot be rejected with a p-value of 0.505. The C matrix of the over-identified MA representation is broadly consistent with the results obtained for the just-identified structure.

As a robustness check, the CRB has been replaced with the Conference Board's Sensitive Materials Index (SENSI). This comprises raw materials and metals but excludes food and energy. The results remain broadly the same as before.

[7] As when imposing just-identifying restrictions on the cointegrating relationships, the PI matrix was uniquely determined, albeit α and β vectors were not.

5. Conclusions

Commodity prices are currently seen as one or the main source of current inflationary pressures and there seems to exist, as discussed by Frenkel (2006), a linkage between increases in commodity prices in commodity exporting countries and monetary policy changes in advanced industrial economies. This, as suggested, would defy the common knowledge that changes in commodity prices are solely impacted by developments occurring in the commodity markets.

The aim of the paper is to identify the nexus between the excess of liquidity in the US and commodity prices over the period 1983-2006 within a cointegrated VAR framework and in particular at testing whether the commodity prices react more powerfully than the consumer goods' prices to changes in real money balances.

The results provide empirical evidence on the magnitude of the reaction of commodity and consumer goods' prices to an increase in real money. In particular, a long-run equilibrium relationship between real money and real commodity prices has been found, with an effect from real output. The two variables of interest show positive and significant correlation.

Moreover, real money seems to be a weakly exogenous variable and therefore is a pushing force (away from the equilibrium), whereas real commodity prices are the only equilibrium correcting variable to such cointegration relationship, or pulling force.

Therefore, in order to restore the long-run equilibrium when there is a real excess (lack) of liquidity, the real commodity prices need to increase (decrease). Such increase (decrease) can be achieved through an increase (decrease) in the commodity prices stronger (smaller) than the increase (decrease) in the consumer goods' prices, generating a larger spread between the two.

The results have important policy implications. More specifically, if the magnitude of the reaction is due to the fact that consumer goods' prices are slower to react, then their long-run value can be predicted with the help of the commodity prices. In other words, the extent of the rise in the commodity prices acts to predict subsequent changes in the price of the other goods, namely the consumer goods, whose price is initially unchanged.

Moreover, the results also support the idea that monetary policy cannot only focus on the core CPI and ignore developments in the commodity market. In fact, if commodity prices are very high it might be the case that monetary policy is loose; therefore they should be taken into account as a useful monetary indicator. This conclusion is particularly relevant to those countries that are adopting an inflation targeting regime which target is the CPI.

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Appendix

The rank test for the I(2) system has been performed for the system of variables (X_t) used in the analysis. The result is border-line as the hypothesis H(3,2,1) cannot be rejected with a p-value of 0.105. Therefore, the system might present rank equal to three, two I(1) trends and one I(2) trends.

However, it should be noted that the estimates obtained are consistent even when there are I(2) trends and that the standard tests (except the I(1) trace test) are still valid.

[Table A about here]

Table 1: Descriptive statistics of the variables

Variable	Obs	Mean	SD	Min	Max
Real Money (M2)	157	7.86	0.72	6.38	9.02
Real GDP	157	24.47	0.27	23.98	24.91
T-Bill Rate	157	5.79	2.92	0.23	15.09
Govt Bond Yield	157	7.38	2.56	2.73	14.85
CPI Inflation	157	4.72	0.51	3.66	5.41
Real Commodity Prices	157	-0.19	0.39	-0.83	0.60

Source: IFS, CRB.

Table 2: Unit Root Tests

Variables	ADF test		KPSS test	
	Level	Difference	Level	Difference
Real Money (M2)	-1.43	-4.80**	1.59**	0.25
Real GDP	-2.98	-4.93**	1.98**	0.28
T-Bill Rate	-3.07	-3.64*	1.25**	0.09
Govt Bond Yield	-3.33	-5.32**	1.70**	0.05
CPI Inflation	-1.57	-3.83*	1.98**	0.45
Real Commodity Prices	-2.73	-6.15**	1.43**	0.17

Notes: The augmented Dickey-Fuller (ADF) test statistics are from a model that includes as many lags as suggested by and the Akaike Information Criterion (AIK) Hannan-Quinn (HQ) criterion, a constant and a trend.

Table 3: Lag Length Determination

Model	k	T	Regr	Log-Lik	SC	H-Q	LM(1)	LM(k)
VAR(5)	5	90	38	3282.75	-61.55	-65.33	0.595	0.856
VAR(4)	4	90	32	3244.50	-62.50	-65.68	0.815	0.016
VAR(3)	3	90	26	3208.81	-63.51	-66.09	0.239	0.139
VAR(2)	2	90	20	3172.93	-64.51	-66.50	0.088	0.731
VAR(1)	1	90	14	3105.83	-64.82	-66.21	0.000	0.000

Notes: k are the lags in the model, T is the number of observations.

Table 4: Multivariate Residual Analysis

Autocorrelation LM(1) LM(2)	Normality	ARCH LM(1) LM(2)
$X^2(36) = 46.019$ [0.122] [0.668]	$X^2(12) = 9.509$ [0.659]	$X^2(441) = 427.608$
$X^2(36) = 37.473$ [0.401] [0.173]		$X^2(882) = 921.552$

Notes: p-values in brackets.

Table 5: Univariate Residual Analysis

Variable	ARCH(2)	Normality	Skewness	Kurtosis
Δ Real Money (M2)	4.717 [0.095]	1.430 [0.489]	0.264	2.736
Δ Real GDP	3.279 [0.194]	4.262 [0.119]	-0.390	3.733
Δ T-Bill Rate	5.583 [0.061]	1.856 [0.395]	-0.215	3.325
Δ Govt Bond Yield	0.787 [0.675]	0.506 [0.777]	0.169	2.868
Δ CPI Inflation	4.800 [0.091]	2.085 [0.352]	0.016	3.411
Δ Real Commodity Prices	1.768 [0.413]	0.458 [0.795]	-0.010	3.050

Notes: p-values in brackets.

Table 6: Trace Test Statistics

$p-r$	r	Eig.Value	Trace	Trace*	Frac95	P-Value	P-Value*
6	0	0.553	190.965	164.447	115.237	0.000	0.000
5	1	0.407	116.182	101.056	85.895	0.000	0.002
4	2	0.264	67.516	59.345	62.538	0.016	0.088
3	3	0.212	38.974	33.960	41.737	0.097	0.254
2	4	0.135	16.859	15.222	25.295	0.406	0.533
1	5	0.035	3.334	2.988	12.486	0.811	0.852

Notes: Bartlett correction of the rank test is denoted by *.

Table 7: α and β vectors for $r = 3$

Vector	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P	Trend
$\hat{\beta}_1'$	6.744	1.000	0.173	0.638	-202.583	0.879	-0.034
$\hat{\beta}_2'$	-3.564	1.000	0.231	-0.237	-8.757	-0.187	0.014
$\hat{\beta}_3'$	-0.310	0.055	-0.101	0.064	1.000	-0.033	0.001

Variables	α_1	α_3	α_3
Real GDP	-0.003	0.008	0.046
	(-1.655)	(0.745)	(2.341)
Real M2	-0.002	-0.019	-0.032
	(-1.223)	(-1.520)	(-1.468)
T-Bill	-0.025	-0.433	0.483
	(-1.049)	(-2.693)	(1.732)
Govt Yield	-0.080	0.577	-0.152
	(-2.677)	(2.809)	(-0.426)
CPI Infl	0.006	0.016	-0.005
	(6.949)	(2.742)	(-0.505)
Real Comm P	-0.016	0.106	0.621
	(-1.274)	(1.220)	(4.118)

Notes: t-values in brackets.

Table 8: Long-Run Exclusion

R	DGF	5% C.V.	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P	Trend
1	1	3.841	1.870	0.464	0.263	2.954	24.229	2.828	2.429
			[0.171]	[0.496]	[0.608]	[0.086]	[0.000]	[0.093]	[0.119]
2	2	5.991	21.386	17.682	13.021	17.659	44.240	6.294	21.047
			[0.000]	[0.000]	[0.001]	[0.000]	[0.000]	[0.043]	[0.000]
3	3	7.815	26.999	17.964	18.735	22.146	50.640	6.388	22.339
			[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.094]	[0.000]
4	4	9.488	29.597	25.850	27.146	23.835	59.199	13.365	22.341
			[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.010]	[0.000]
5	5	11.070	38.923	30.867	37.155	33.683	68.021	20.884	30.340
			[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.001]	[0.000]

Notes: p-values in brackets.

Table 9: Test for Weak Exogeneity

R	DGF	5% C.V.	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P
1	1	3.841	2.071	1.213	0.710	4.539	19.275	0.998
			[0.150]	[0.271]	[0.400]	[0.033]	[0.000]	[0.318]
2	2	5.991	2.442	3.112	5.357	11.204	35.321	1.709
			[0.295]	[0.211]	[0.069]	[0.004]	[0.000]	[0.425]
3	3	7.815	4.538	4.884	6.952	11.345	36.193	6.640
			[0.209]	[0.181]	[0.073]	[0.010]	[0.000]	[0.084]
4	4	9.488	10.855	4.988	9.042	12.632	36.895	12.794
			[0.028]	[0.289]	[0.060]	[0.013]	[0.000]	[0.012]
5	5	11.070	14.584	7.949	18.947	20.752	46.093	20.846
			[0.012]	[0.159]	[0.002]	[0.001]	[0.000]	[0.001]

Notes: p-values in brackets.

Table 10: Test for Pure Adjustment

R	DGF	5% C.V.	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P
1	5	11.070	44.672	53.863	33.205	27.097	7.898	49.811
			[0.000]	[0.000]	[0.000]	[0.000]	[0.162]	[0.000]
2	4	9.488	23.456	31.534	8.286	11.106	7.337	24.816
			[0.000]	[0.000]	[0.082]	[0.025]	[0.119]	[0.000]
3	3	7.815	6.450	17.230	7.770	4.286	7.311	5.124
			[0.092]	[0.001]	[0.051]	[0.232]	[0.063]	[0.163]
4	2	5.991	1.927	13.706	4.685	4.136	6.926	1.227
			[0.382]	[0.001]	[0.096]	[0.126]	[0.031]	[0.541]
5	1	3.841	0.421	10.151	0.318	0.541	4.807	1.064
			[0.517]	[0.001]	[0.573]	[0.462]	[0.028]	[0.302]

Notes: p-values in brackets.

Table 11: The just-identified long-run cointegration relations for $r = 3$ and the α coefficients.

Vector	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P	Trend
$\hat{\beta}_1'$	-0.979	1	-0.598	0.598	-62.152	0	-0.001
	(-1.796)		(-6.575)	(6.575)	(-11.789)		(-0.171)
$\hat{\beta}_2'$	-4.148	1	-0.145	0	0	-0.285	0.016
	(-10.745)		(-4.319)			(-4.279)	(6.391)
$\hat{\beta}_3'$	0	0	0.170	-0.129	1	0.017	0.001
			(9.357)	(-6.876)		(0.870)	(1.686)

Variables	α_1	α_3	α_3
Real GDP	-0.009	0.017	-0.036
	(-1.488)	(1.691)	(-0.949)
Real M2	-0.010	-0.013	-0.056
	(-1.543)	(-1.123)	(-1.355)
T-Bill	-0.177	-0.003	-0.017
	(-2.055)	(-1.756)	(-3.258)
Govt Yield	-0.165	0.007	0.008
	(-1.500)	(3.533)	(1.127)
CPI Infl	0.023	-0.002	0.111
	(7.484)	(-0.335)	(5.756)
Real Comm P	-0.052	0.176	-0.275
	(-1.118)	(2.248)	(-0.952)

Notes: t-values in brackets.

Table 12: The over-identified long-run cointegration relations for $r = 3$ and the α coefficients.

Vector	Real GDP	Real M2	T-Bill	Govt Yield	CPI Infl	Real Comm P	Trend
$\hat{\beta}_1'$	-1	1	4.045	-3.832	46.431	0	0
			(9.401)	(-8.880)	(10.579)		
$\hat{\beta}_2'$	-4.073	1	0	0	0	-0.338	0.018
	(-11.655)					(-6.603)	(8.278)
$\hat{\beta}_3'$	0	0	0.039	-0.039	1	0	-0.000
			(9.041)	(-9.041)			(-2.998)

Variables	α_1	α_3	α_3
Real GDP	-0.012	0.018	1.114
	(-1.753)	(1.615)	(1.699)
Real M2	-0.011	-0.010	1.078
	(-1.459)	(-0.808)	(1.489)
T-Bill	-0.242	-0.365	20.742
	(-2.580)	(-2.341)	(2.265)
Govt Yield	-0.153	0.535	19.056
	(-1.241)	(2.606)	(1.577)
CPI Infl	0.025	-0.004	-2.505
	(7.439)	(-0.746)	(-7.494)
Real Comm P	-0.062	0.208	5.581
	(-1.187)	(2.392)	(1.093)

Notes: t-values in brackets.

Table 13: Short-run Identification, equation for the real commodity prices

Independent Variable	Coefficient	Std.Error	t-value	t-prob
Second Coint Rel	0.166	0.060	2.75	0.007
Constant	15.784	5.738	2.75	0.007
CSeasonal	-0.006	0.013	-0.434	0.666
CSeasonal_1	-0.004	0.013	-0.322	0.748
CSeasonal_2	-0.009	0.013	-0.656	0.514

Table 14: Correlation of Structural Residuals (standard deviations on diagonal)

−	Real GDP	*Real M*2	*T* − *Bill*	Govt Yield	CPI Infl	Real Comm P
Real GDP	0.006	−	−	−	−	−
*Real M*2	0.463	0.006	−	−	−	−
T − *Bill*	0.209	−0.350	0.001	−	−	−
Govt Yield	0.290	−0.149	0.664	0.001	−	−
CPI Infl	−0.443	−0.601	0.219	0.318	0.003	−
Real Comm P	0.226	−0.127	0.420	0.374	0.193	0.045

Table A: I(2) Rank Test

p-r	r	s2 = p-r-s1 6	5	4	3	2	1	0
6	0	585.487 (0.000)	430.095 (0.000)	349.873 (0.000)	283.396 (0.000)	241.097 (0.000)	209.188 (0.000)	190.965 (0.000)
5	1		340.790 (0.000)	270.370 (0.000)	207.163 (0.000)	167.098 (0.000)	134.455 (0.000)	116.182 (0.000)
4	2			207.162 (0.000)	152.244 (0.000)	108.296 (0.003)	81.881 (0.018)	67.516 (0.022)
3	3				102.274 (0.003)	66.687 (0.084)	47.417 (0.105)	38.974 (0.117)
2	4					40.874 (0.239)	26.709 (0.316)	16.859 (0.434)
1	5						12.445 (0.424)	3.334 (0.827)
Approximate 95% Fractiles								
6	0	282.595	244.789	211.074	181.464	155.978	134.640	117.451
5	1		206.055	174.292	146.636	123.112	103.747	88.554
4	2			141.531	115.818	94.243	76.841	63.659
3	3				89.020	69.376	53.921	42.770
2	4					48.520	34.984	25.731
1	5						20.018	12.448

Notes: p-values in brackets. The critical values correspond to the “Basic Model”.

Figure 1: Short and long-run impact of a liquidity shock to price elastic and price inelastic good.

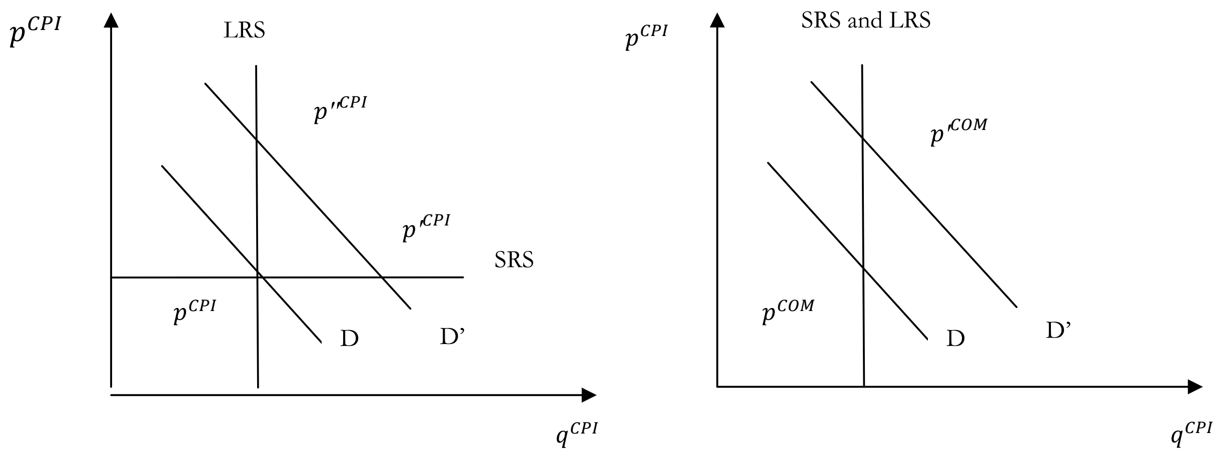

Figure 2: Data in levels

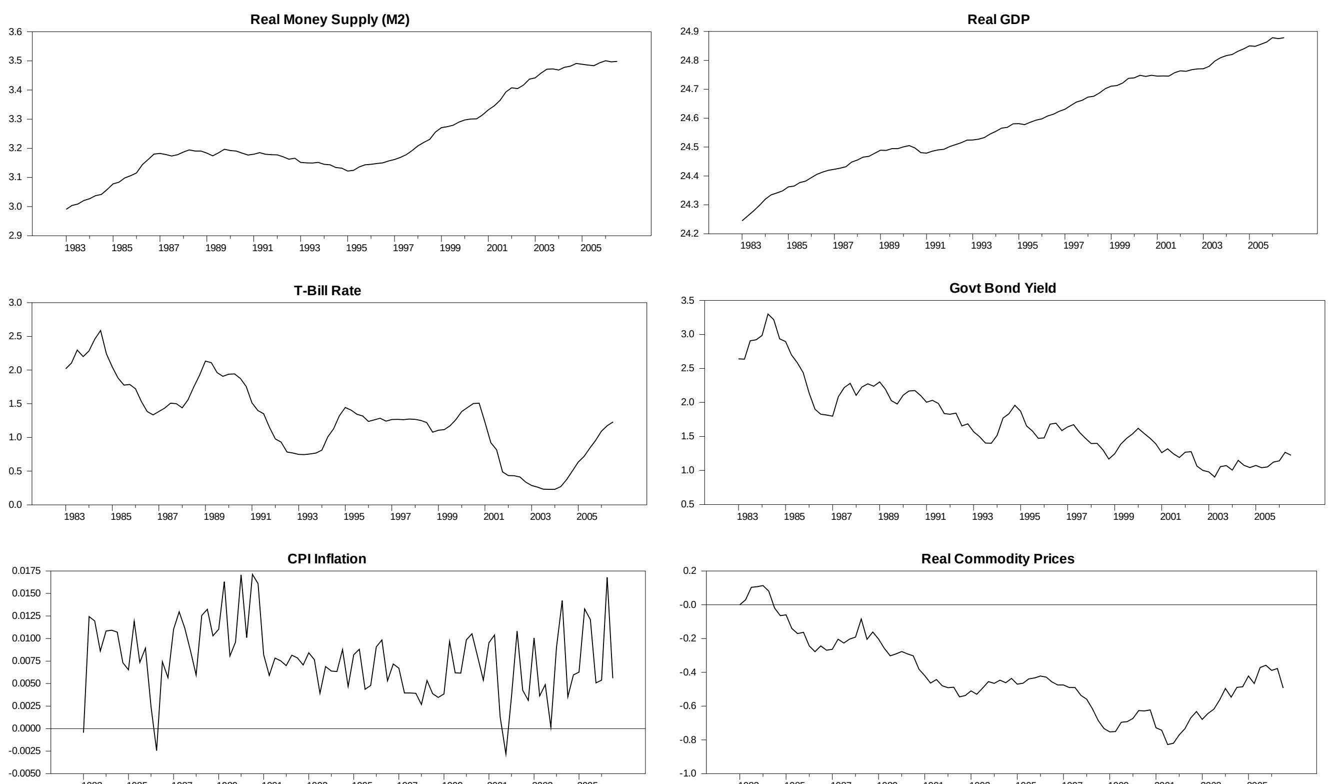

Figure 3: Data in first differences

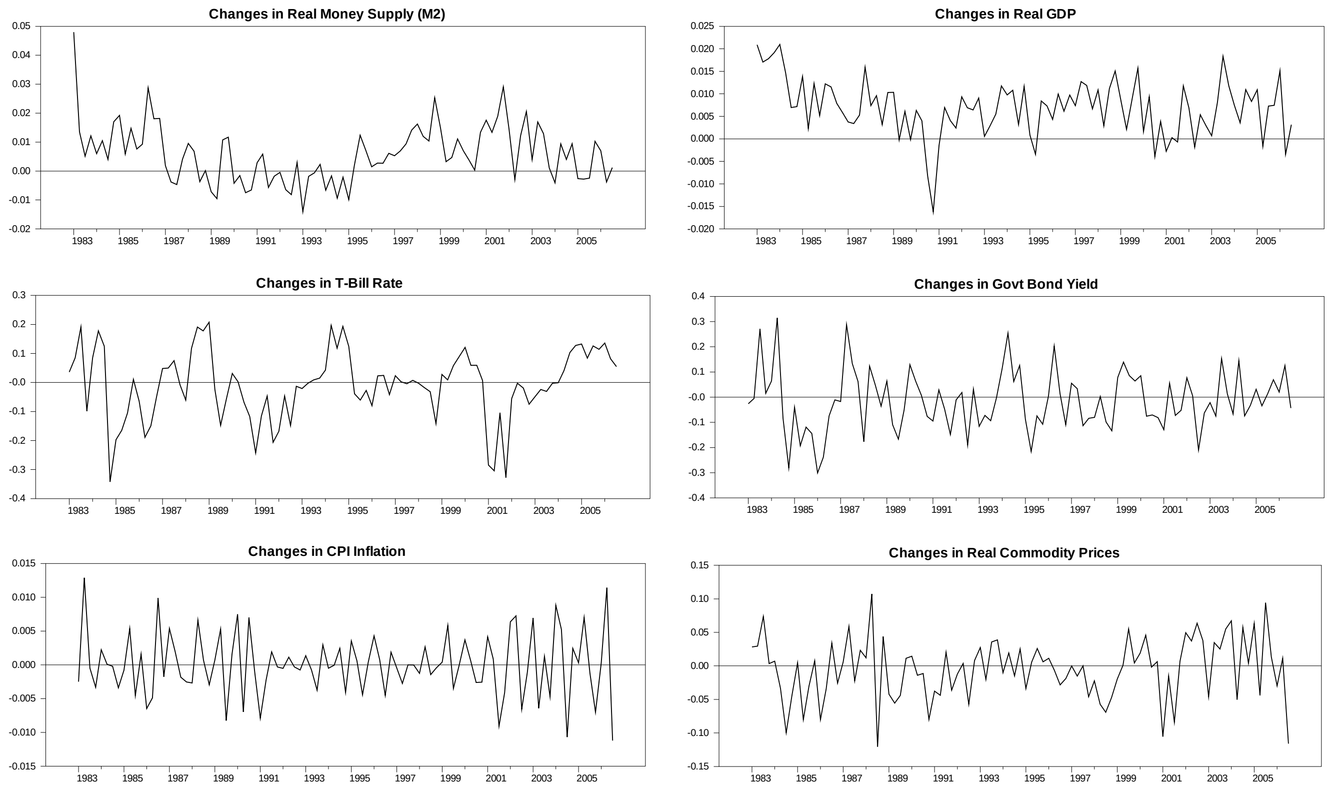

Figure 4: Cointegrating relationships

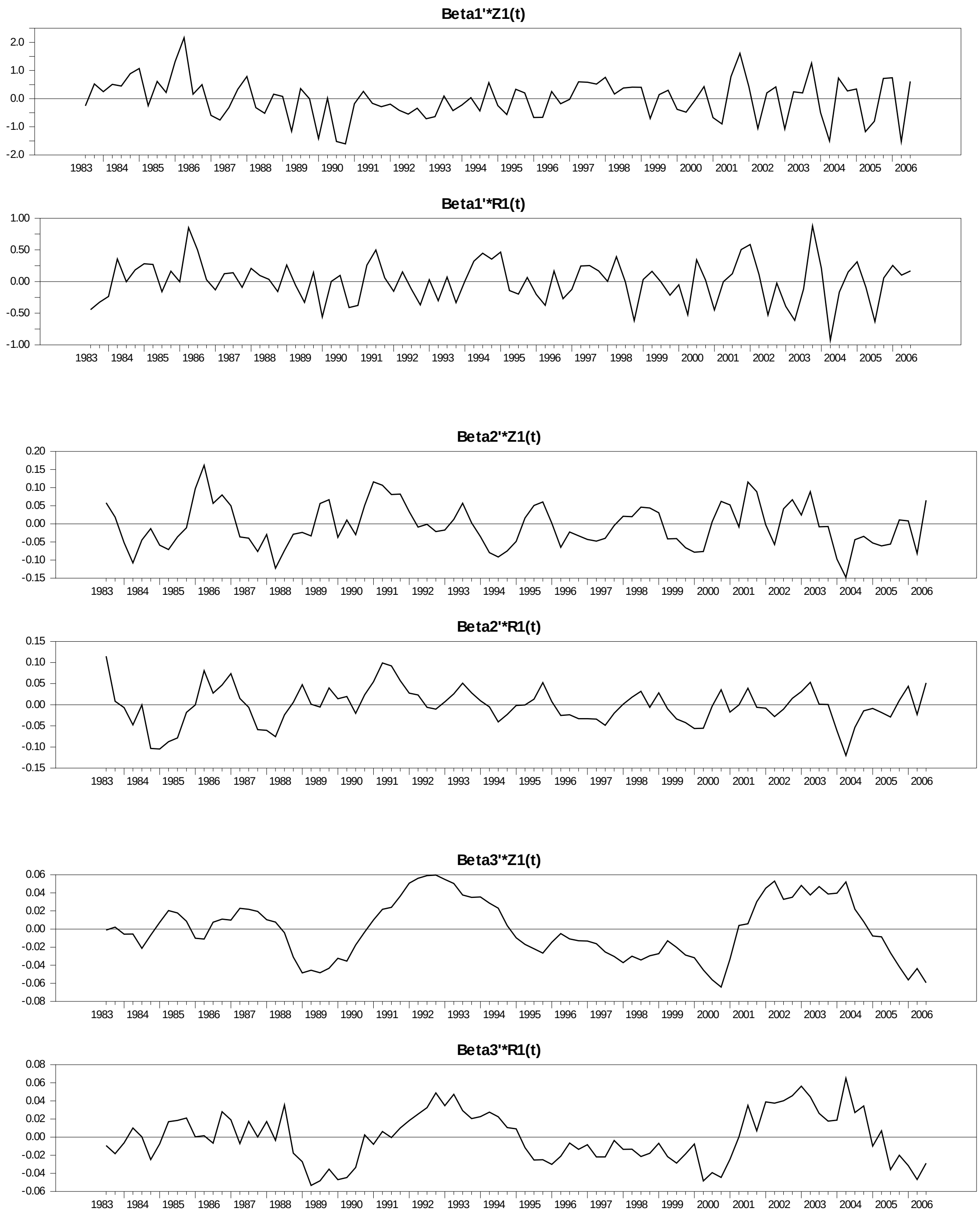

Figure5: Roots of the Companion Matrix for $r = 3$

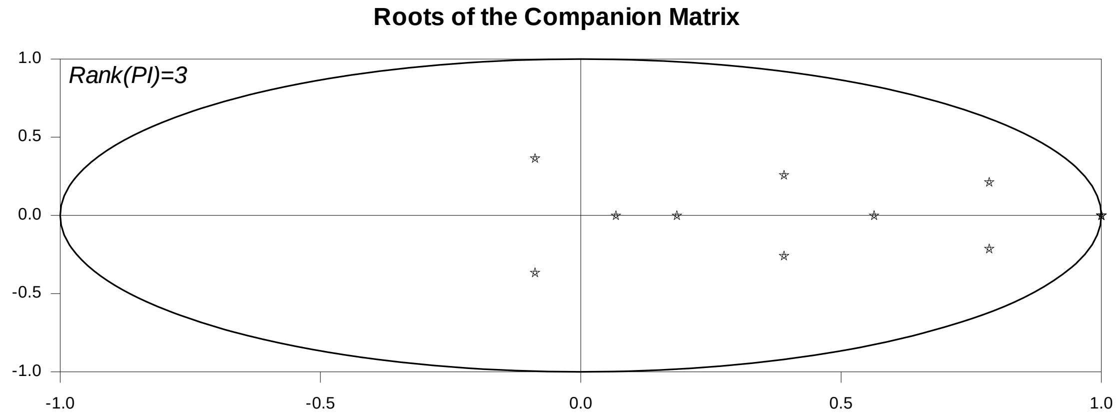

Figure 6: Eigenvalue Fluctuation Test for $r = 3$

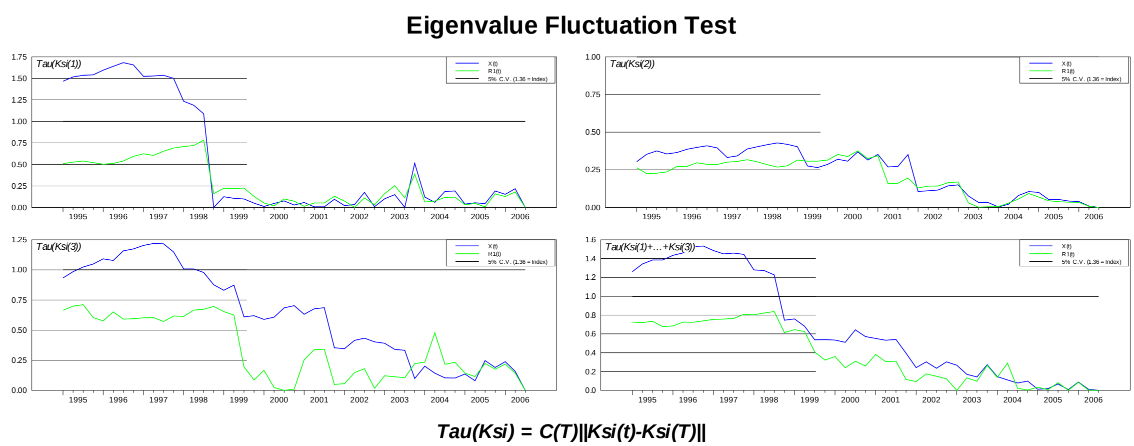

Figure 7: Max Test of Constant β

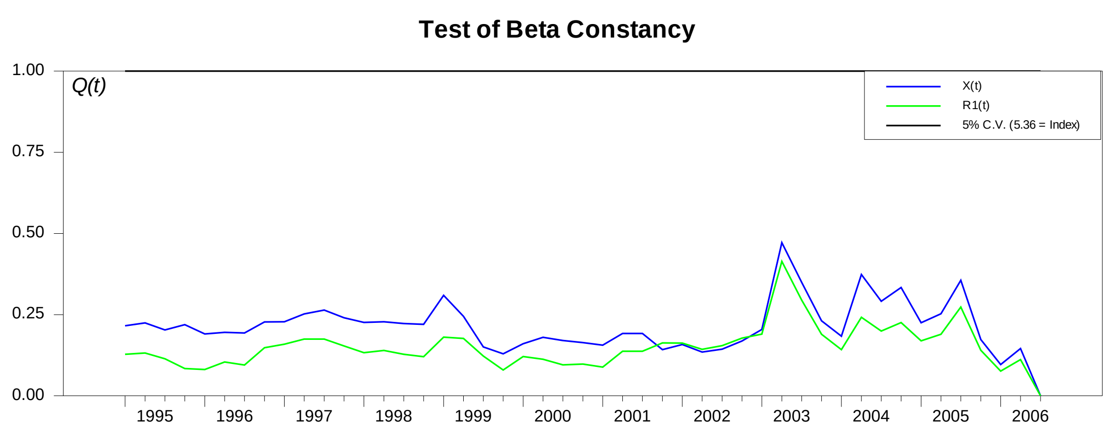

Figure 8: Commodity Prices and Consumer Goods´ prices reactions to Real Excess of Liquidity